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# THE DEFAULT RISK OF HIGH-YIELD BONDS

by

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THE DEFAULT RISK OF HIGH-YIELD BONDS

Abstract

This paper investigates the default behavior of original issue rated non-convertible high-yield bonds. The hazard model simultaneously estimates the impact of bond age, firm- and issue-specific characteristics, and changing economic conditions. The specification used models the impact of the time since issuance semi-parametrically, corrects for unobserved heterogeneity and allows for the possibility that outstanding bonds may default in the future. Our findings, based on a sample of 579 individual high-yield bonds issued between 1977 and 1989, suggest that, after controlling for annual changes in economic conditions, default rates increase with age. Bond characteristics at the time of issuance also impact the default behavior: BB rated bonds tend to have significantly lower default rates compared to CCC rated bonds; bonds with higher coupon rates have a significantly higher default rates. In addition, high-yield bonds issued prior to 1980 experienced significantly lower default rates.

J.E.L. Classifications: G19, G33, C41

## I. Introduction

During the late 1980s, the market for high-yielding, low-rated corporate bonds, or "junk bonds," grew rapidly to about 25% of all public domestic corporate issues outstanding by the end of 1993 [Moody's (1994)]. Given the increase in the new issue volume of high-yield corporate debt, and the high default rates in 1990 and 1991, numerous concerns have been expressed by investors and policymakers regarding the default risk within this market. This study investigates empirically the determinants of defaults of high-yield bonds.

As the market for high-yield corporate bonds matures, one issue of concern is whether default risk changes as bonds age. Many studies note that the probability of default for newly issued high-yield bonds is initially low and then dramatically increases over the first two to four years after issuance.[1] Low default rates immediately subsequent to issuance suggest companies avoid issuing bonds with high initial default risk. However, it is not clear that high-yield bonds will become more or less likely to default once they pass this critical point. A bond's survival for the first few years after issuance may provide positive information regarding the issuer's ability to manage its debt burden. Alternatively, a bond that remains outstanding may reflect that the creditworthiness of the issuer has not improved sufficiently to call the bond and refinance at a lower rate.

The empirical evidence on the relationship between default risk and bond age is mixed. Altman (1989) and Asquith, Mullins and Wolff (1989), for example, found that the longer bonds have been outstanding, the lower the default probability. However, Blume, Keim and Patel (1991) proposed that this aging relationship may in fact reflect changing economic conditions. We address this concern by examining simultaneously the effect of aging and the impact of macroeconomic changes.

In addition to macroeconomic conditions and age, default rates may also be related to issue-specific attributes. Characteristics which have been shown to be associated with default of high-yield bonds include: the bond's rating at issuance, the coupon rate, the seniority, the year of issuance, the bond's maturity and whether the bond was underwritten by Drexel [see Altman (1992); Asquith et al. (1989); Cheung, Bencivenga and Fabozzi (1992); Cotter and Peck (1995); Platt (1993); and Rosengren (1993)].

Much of this existing research (a) used aggregate analyses, and/or (b) studied the potential influence of a particular bond characteristic in isolation.[2] In contrast, this paper employs individual-level survival analysis to investigate simultaneously the impact of age, changing economic conditions and issue-specific characteristics on the default rate. We use a flexible hazard model specification where the impact of the time since issuance is modeled semi-parametrically, thereby imposing no distribution on the estimation of the aging effect. Furthermore, the chosen specification takes unobserved heterogeneity into account; we show how a failure to do so affects the estimates. By taking into account both observed (through the explanatory variables) and unobserved heterogeneity, the hazard model used in this study allows us to separate spurious from real time dependence.

The model is applied to a data set of 579 individual high-yield non-convertible corporate bonds issued between 1977 and 1989. We find that, after controlling for annual changes in economic conditions, default rates increase with bond age. Bonds rated BB at issuance have significantly lower default rates than CCC bonds. When the rating, age and economic conditions are taken into account, other issue characteristics such as seniority, initial maturity, whether the bond was underwritten by Drexel, and whether the bond was issued as part of a leveraged buyout (LBO) do not appear to have a significant influence. The coupon rate, and the year of issuance, however, are significantly related to the default probability.

The remainder of the paper is organized as follows. Section II reviews the empirical literature on the determinants of bond defaults and distinguishes our study from prior work. The flexible hazard specification is developed in Section III. Section IV describes the sample and explanatory variables. Empirical results are presented in Section V. Finally, Section VI summarizes the findings.

## II. Determinants of High-Yield Bond Defaults

***Aging Effect:*** Altman (1989) and Asquith, et al. (1989) examined the high-yield corporate bond market and presented evidence that bonds outstanding for longer periods of time are more likely to default, which is referred to as an "aging effect". Support for the aging effect has often been based on an actuarial approach: aggregate default rates for bonds that have been outstanding for equal periods of time (i.e., all issued in the same cohort year),[3] were calculated and shown to increase over time.

***Macroeconomic Conditions:*** Blume, Keim and Patel (1991) questioned the presence of an aging effect by observing that in some years there are more defaults than in others.[4] For example, default rates were high across all age groups in 1985, 1987, 1990 and 1991, and were low across all age groups in 1988. Blume et al. (1991) thus proposed that a large portion of the defaults previously attributed to bond age may be more appropriately attributed to overall economic conditions. Indeed, Blume and Keim (1991) found that the aging effect disappeared in their sample when default rates were adjusted for overall economic conditions.

Macroeconomic conditions may affect the default probability in a number of ways. First, as argued in Denis and Denis (1995), recessions such as 1990-1991 are likely to reduce cash flows, thereby increasing the likelihood that issuers cannot meet their obligations. Second, Shleifer and Vishny (1992) argued that asset markets are less liquid during an economic downturn, which may deter firms from selling assets in order to fund their debt obligations. The approach taken in this study allows us to distinguish between the aging effect and the impact of macroeconomic changes.

Another problem when making inferences about the presence of an aging effect from aggregated data is that the observed pattern may result from spurious aggregation effects. Lancaster (1990), Morrison and Schmittlein (1980) and Schmittlein and Morrison (1983) demonstrated that aggregation may result in downward biased estimates of the time dependence. This bias can be illustrated by considering a situation where none of the bonds experience any aging effect, but where some bonds have a high default rate, while others are characterized by an extremely low default rate. The high-default-rate bonds will default early on, leaving a higher proportion of bonds with a very low default probability. This will cause the observed number of defaults to become smaller, which would suggest a negative correlation between default and age, even without any individual bond's default probability changing with age. The hazard model used in this study allows us to separate spurious from real time dependence by taking into account both observed and unobserved heterogeneity.

***Rating:*** Using the actuarial approach as described above, i.e., where the aggregate default rate is calculated for specific cohort groups, prior studies have found evidence that the default rate of high-yield bonds is related to the rating provided by Moody's and Standard & Poor's at issuance. For example, Altman (1992) and Asquith et al. (1989) found that B rated issues tended to exhibit higher cumulative default rates than BB rated issues. Similarly, Altman and Kishore (1995) found that CCC rated issues experienced higher cumulative default rates compared to BB and B; Hradsky and Long (1989) reported a decline in the average time from issuance to default from higher- to lower-rated bonds. However, it is unlikely that ratings capture all relevant individual-level factors which may affect the likelihood of default.

***Coupon Rate:*** In addition to the bond's rating, the coupon rate may provide information about the bond's default risk. A company that is committed to paying a higher fixed interest rate is likely to have a higher probability of defaulting on the payments, all other things equal. Lehman and Fridson (1995), for example, found that even after normalizing for ratings, high coupon issues are more likely to default than low coupon issues.

***Seniority:*** Cheung, Bencivenga and Fabozzi (1992) found that seniority is related to bond default. Alternatively, Fridson and Garman (1995) argued that while subordinated bonds are expected to have lower payoffs in the default state compared to senior bonds with the same rating, the former are not expected to have a higher default probability.

***Underwriter (Drexel):*** Prior studies have found that a smaller percentage of bonds underwritten by Drexel Burnham Lambert experienced default compared to bonds with other underwriters [Altman (1989); Asquith et al. (1989); Platt (1993)]. This underwriter effect has been attributed to Drexel's dominance, experience and expertise in the high-yield market and attempts by other underwriters to penetrate the market by underwriting issues of less creditworthy firms.

***LBO-Related:*** Whether the issue was used to finance a leveraged buyout may also contribute to the default risk. LBO-related issues, for example, may be more risky than other newly-issued high-yield debt because LBO firms tend to be more highly leveraged. Wigmore (1990) provided a simple model of LBO returns which suggests that for most issuers, debt reduction needs to be combined with growth in earnings if long-term debt obligations are to be met. He noted that while most LBOs create cash by selling assets and cutting capital expenditures in the early years, this may hinder the sustained growth needed for the next 10 to 15 years if the issuer is to be able to repay debt.

***Maturity Structure:*** The maturity structure of the debt may contribute to its default risk. Cotter and Peck (1995) hypothesized that shorter maturity debt is associated with a greater likelihood of default. They argued that issuing shorter maturity debt increases the debt burden in early periods which, in turn, increases the probability of default. Cotter and Peck examined a sample of LBOs, and provided evidence that as the average maturity (at issuance) of the debt increased, the likelihood of subsequent default decreased.

***Issuance Year:*** Wigmore (1990) provided evidence of a decline in credit quality for high-yield bonds issued in the 1980's. He observed that the decline in credit quality was only partially reflected by increases in the percentage of bonds issued with the lowest credit ratings. He showed that credit quality (based on financial ratios) within rating categories declined after 1985.

In most studies, each of the above factors, including the aging effect, has been considered in isolation. An exception is Altman (1989), where the cohorts consisted of bonds issued in a given year and in a given rating. Clearly, such an actuarial procedure becomes cumbersome when dealing with multiple predictors, and the number of observations in each cell will reduce quickly. Rosengren (1993) adopted a logit model to estimate the probability of default as a function of bond-specific features. Consistent with default rates increasing with bond age, Rosengren found that the probability of default is related to the year of issuance, with earlier years having significantly higher default rates than later years. Drawbacks of this approach, as illustrated in Gupta (1991), however, are that (1) bonds which are still outstanding at the end of the observation period are classified as non-defaulting even though they may default in the future, (2) for those bonds which have defaulted, no distinction is made between early and late defaults, which results in a loss of information, and

(3) it is difficult to correct for unobserved heterogeneity and to incorporate time-varying explanatory variables in logit models.[5]

Our work extends the empirical literature on the individual-level default risk of high-yield bonds in several ways. First, we investigate the impact of aging, changing economic conditions and issue-specific characteristics simultaneously. Second, we use a flexible hazard model specification that (a) models the impact of aging semi-parametrically, which allows for a monotonic or non-monotonic aging effect, (b) includes both fixed and time-varying variables, and (c) accounts for unobserved heterogeneity. Lastly, our observation period ends in December 1994, which is a longer observation period than prior work.

## III. Method of Analysis

Let the random variable $T$ denote the time between issuance and default, with associated probability density function $f(t)$ and cumulative density function $F(t)$. Then,

$$h(t) = \frac{f(t)}{1 - F(t)} , \tag{1}$$

is the hazard, i.e., the rate at which bonds default during period $t$ given that they have not done so in the previous $t$-1 periods since issuance. The survival function $S(t)=1-F(t)$ denotes the probability that default does not occur for at least $t$ periods. To account for the discrete nature of our data, monthly grouping intervals $[t_{k-1}, t_k)$, $k=1,2,...,m+1$, $t_0=0$ and $t_{m+1}=\infty$ are defined, and default in interval $[t_{k-1}, t_k)$ is recorded as $t_k$, which is the number of months since issuance.

As indicated before, the default rate of bonds may depend not only on the time elapsed since issuance, but also on the bond's characteristics and on general economic conditions. To model the hazard rate as a function of explanatory variables, we use the Cox (1972) formulation and let

$$h(t) = h_0 \, e^{\beta X(t)} \;, \tag{2}$$

where $X(t)$ is a vector of explanatory variables which may be fixed (e.g., coupon rate) or time-varying (e.g., economic conditions), $\beta$ is a vector of parameters, and $h_0$ is the default rate of the base group. The base group consists of bonds for which all explanatory variables equal zero. A positive $\beta$-coefficient implies that a positive value of the associated variable augments the default rate. Specifically, when the $j$-th variable changes by one unit, the hazard changes by $[\exp(\beta_j)-1)]*100$ percent.

To quantify the impact of the time since issuance, one can either assume a particular distribution for the time dependence of default rates, or introduce this time dependence in a semi-parametric fashion. Given the absence of a theoretical justification for choosing a particular distribution, the latter approach is preferred, especially since the choice of an incorrect parametric distribution has been shown to result in inconsistent parameter estimates. The semi-parametric approach that we adopt, on the other hand, results in consistent estimates even when the true form of the underlying baseline hazard is unknown, as shown in Meyer (1986,1990). The semi-parametric approach involves adding a vector of time-varying dummy variables, $D(t)$, to the model:[6]

$$h(t) = h_0 \, e^{\beta \, X(t)} \, e^{c \, D(t)} \;, \tag{3}$$

where $c$ is a vector of coefficients. A separate variable is used for each period; for example, D(3) takes on the values (0 0 1 0 0 ...).[7] The quantity $h_0$ in equation (3) then gives the default rate of the base group in the first post-issuance period. Positive (negative) $c$-coefficients indicate a higher (lower) default rate compared to the first period.

The semi-parametric approach makes no distributional assumptions regarding the nature of the time dependence.. The only assumption made is that within a grouping interval, the hazard remains constant. The aging effect is thus measured as a piecewise approximation of an underlying, possibly very complex, continuous time-dependence pattern. If the default rate is constant over time (i.e., there is no aging effect), all the coefficients of the time-varying dummy variables are equal to zero and $t$ follows an exponential distribution. Monotonically increasing $c$-coefficients imply that the default probability increases as the bond ages. However, our model is flexible enough to also allow for a non-monotonic dependence on time.

Several patterns of time dependence may be possible for high-yield bonds. The studies cited in Section II suggest a positive time dependence. Still, one could also argue that a bond's age provides information about the issuer's ability to manage its debt burden. Put differently, a bond remaining outstanding for a large number of periods may indicate that the issuer has been able to meet its debt payments for a long time. The (conditional) likelihood of default may therefore start to decrease from a certain point onwards, in which case a non-monotonic pattern would be observed.

To estimate the parameters $h_0$, $c$ and $\beta$, maximum likelihood estimation is used. In what follows, we first derive an expression for the likelihood function in terms of the survival function, after which we use a general relationship between a distribution's hazard and survival functions to express the likelihood function in terms of $h_0$, $c$ and $\beta$.

The contribution of a given bond to the likelihood function depends on whether the bond is classified as a completed or censored observation. In practice, one can distinguish between five types of non-convertible bonds: (1) bonds that defaulted, (2) bonds that were still outstanding at the end of the observation period, (3) bonds that were called, (4) bonds that matured, and (5) bonds that were exchanged. The first category can be interpreted as completed observations, while the other four categories are referred to as censored.

A bond $X$ that has defaulted is a completed observation: we know the time $t_X$ that has elapsed since issuance before the bond defaulted. Because of the discrete nature of the data (i.e., we are working with monthly grouping intervals), the contribution of a defaulted bond to the likelihood function is given by the difference in the survival functions $S(t_X\text{-}1)\text{-}S(t_X)$. Bonds that are still outstanding by the end of the observation period are censored observations: these bonds may or may not default in the future. The contribution to the likelihood function of such a bond $Y$ with an observed duration of $t_Y$ is given by the survival function $S(t_Y\text{-}1)$: we know the bond has not defaulted within the first $t_Y$-1 months.[8] Bonds that have been called are also classified as censored; since these bonds can no longer default after being called, their hazard becomes zero at that point (or, alternatively, their survival function remains constant after the call date). Hence, the contribution to the likelihood function of a bond $C$ that has been called is given by the survival function evaluated at the number of months prior to the call or exchange date $S(t_C\text{-}1)$. Using similar reasoning, a bond $M$ that has matured is also censored since this bond can no longer default after reaching maturity; its contribution to the likelihood function is therefore given by $S(t_M)$, where $M$ is the number of months in the life of the bond. For our purposes, bonds involved in non-distressed exchanges are also treated as censored observations. However, as in Platt (1993), we classify distressed exchanges as defaults. Based on the above arguments, the contribution of the $i$-th bond to the likelihood function is given by the following expression:

$$L_i(t_i \mid h_0) = [S_i(t_i - 1 \mid h_0) - S_i(t_i \mid h_0)]^{1-d_i} \, [S_i(t_i - 1 \mid h_0)]^{d_i} \,, \tag{4}$$

where $t_i$ is the number of observed periods (months) until default or censoring, and $d_i$ is an indicator variable equal to one for censored observations and zero for completed observations.[9]

The expression for the hazard function given in equation (3) can be substituted into equation (4) through the following general relationship [Lancaster (1990); Ross (1980)]:

$$S_i(t_i) = e^{-\theta_i(t_i)} \text{ where } \theta_i(t_i) = \int_0^{t_i} h_i(u) \, du \,. \tag{5}$$

If we assume that the time-varying variables remain constant within each period but that they can change from period to period, $\theta_i(t_i)$ can be written as:

$$\begin{aligned} \theta_i(t_i) &= \int_0^1 h_i(1) du + \int_1^2 h_i(2) du + \ldots + \int_{t_i - 1}^{t_i} h_i(t_i) du \\ &= h_i(1) + h_i(2) + \ldots + h_i(t_i) \\ &= h_0 \, e^{\beta X_i(1) + c D_i(1)} + h_0 \, e^{\beta X_i(2) + c D_i(2)} + \ldots + h_0 \, e^{\beta X_i(t_i) + c D_i(t_i)} \\ &= h_0 \, B_i(t_i) \,, \text{ where } B_i(t_i) = \sum_{j=1}^{t_i} e^{\beta X_i(j) + c D_i(j)} \,. \end{aligned} \tag{6}$$

After appropriate substitutions, it can be shown that the log-likelihood function for $N$ bonds is:

$$LL(t_i \mid h_0) = \sum_{i=1}^{N} \{ (1-d_i) \ln[e^{-h_0 B_i(t_i-1)} - e^{-h_0 B_i(t_i)}] - d_i h_0 B_i(t_i-1) \} . \tag{7}$$

While the explanatory variables may capture some of the cross-sectional variation, part of the heterogeneity may not be quantifiable. For example, it would be difficult to measure the investment opportunity set of the issuer. Yet changes in this opportunity set would affect the ability of the issuer to meet its debt commitment. The presence of unobserved heterogeneity leads to a downward bias in the $c$-estimates (which would translate into an underestimation of the aging effect), and to inconsistent estimates of the $\beta$-coefficients [Heckman and Singer (1984a,b)]. Unobserved heterogeneity can be modeled by allowing $h_0$ to vary across bonds according to a certain distribution. In other words, we allow for bonds with the same observable explanatory variables to have a different mean default probability. Mathematically, this can be accomplished by weighing the conditional likelihood by the relative occurrence of its $h_0$-value. As such, one obtains the unconditional likelihood contribution of the $i$-th bond, which no longer depends on a specific value of $h_0$:

$$L_i(t_i) = \int_0^{\infty} L_i(t_i \mid h_0)\, g(h_0)\, dh_0 , \tag{8}$$

where $g(h_0)$ is called the unobservable mixing distribution. A commonly used mixing distribution is the gamma distribution [see e.g. Meyer (1990)]. The gamma mixing distribution is often used because of its flexibility (it can take on both inverted U and J shapes) and because it gives a closed-form solution for the log-likelihood function:[10]

$$LL = \sum_{i=1}^{N} \ln \{ (1+d_i) [\frac{a}{B_i(t_i-1) + a}]^r - [\frac{a}{B_i(t_i-1) + (1-d_i)\, e^{\beta X_i(t_i) + c D_i(t_i)} + a}]^r \} . \tag{9}$$

The mean of the gamma distribution, $r/a$, is then used as an estimate of the hazard of the base group in the first period, and the $\beta$ and $c$- parameters are subsequently interpreted relative to this mean in the same way as they were interpreted relative to $h_0$.

## IV. Data

The data set consists of 1185 original-issue high-yield non-convertible corporate bonds identified by Securities Data Corporation as publicly issued in the United States between January 1977 and December 1989. We limit the sample to the 703 bonds for which the issuer is included in the *Compustat* files; i.e., those bonds for which post-issuance information is more likely to be available. The rating, seniority, and underwriter of each bond are identified for 579 bonds with issuance information in the Standard & Poor's *Monthly Bond Guides*. Using the same source, we identify defaults, calls, exchanges, and bonds that were still outstanding as of December 1994.[11] Default is defined as the assignment of a D rating by S&P for a missed coupon payment. Our definition of default is similar to that of other studies, including Asquith et al. (1989), Blume and Keim (1991), and Rosengren (1993). For those bonds where the post-issuance history was not apparent from this source, the Moody's *Bond Record*, the *Wall Street Journal* and *Lexis/Nexis* were consulted. For these bonds, default is defined as a report of a missed coupon, or bankruptcy filing. The sample includes three distressed and three non-distressed exchanges.

Tables 1 and 2 present summary statistics on the issue-specific explanatory variables. Table 1 describes the sample per year of issuance. For our sample, the number of high-yield bonds issued and the average size increased steadily through the 1980s. These observations suggest that our sample is a representative sample of new issuances of high-yield bonds in 1977-1989 [see e.g., Altman (1992), Drexel Burnham Lambert (1989)]. As of December 1994, 27.6% of all issues had defaulted. The proportion of defaults is higher than that of prior studies, presumably because of the longer observation period we use here. The mean annual coupon rate varied from a low of 10.47% in 1977 to a high of 13.74% in 1980.

Table 1 also reports the number of bonds issued in each rating category by year of issuance.[12] The table illustrates the decline in credit quality of new-issue junk bonds: there is a dramatic increase in the percentage of high-yield new issuances rated in the lower rating categories in the late 1980s.

Table 2 provides the number of bonds issued, defaulted, and mean time to default, for each initial rating, each seniority, Drexel underwritten bonds and LBO-related bonds. The majority of the bonds were rated B at issuance, with only a small proportion of bonds rated CCC. Overall, B and CCC rated bonds experienced a larger default percentage (29.0% and 29.1%, respectively) compared to BB rated bonds (19%). Of defaulting bonds, the mean time to default was 4.54 years. Bonds initially rated CCC had the shortest mean time to default (4.10 years). For censored bonds, the mean time until censored was 6.85 years.

The majority of bonds issued were subordinated or senior subordinated. The number of bonds defaulting in our sample is greatest for subordinated debt (29.4%). Of bonds which defaulted, senior subordinated had the shortest mean time to default (4.00 years).

The proportion of bonds underwritten by Drexel is similar to that in other studies; of the 579 bonds included in our sample, 230 are underwritten by Drexel. While the proportion of bonds underwritten by Drexel which default (30.4%) is slightly higher than that of the sample (27.6%), the mean time to default is the same for Drexel and the overall sample (4.54 years).

The bonds in our sample that were associated with LBOs were identified by Securities Data Corporation. *Lexis/Nexis* was used to verify this information. Sixty of the bonds in the sample were associated with LBOs. Fourteen of the sixty LBO-related bonds, or 23.3%, defaulted, which is lower than the default rate of 27.6% for the remainder of the sample and the 27% default rate for management buyouts reported in Kaplan and Stein (1993).

As a measure of overall economic conditions throughout the 1977-1994 period, the change in the Gross National Product (GNP) is included as a time-varying explanatory variable. For each month *t* subsequent to a bond's issuance, the annual percentage change in GNP is calculated using GNP aggregated over the preceding 12 months, i.e. $GNP_{t-1}$ through $GNP_{t-12}$, relative to $GNP_{t-13}$ through $GNP_{t-24}$. The probability of default should be negatively related to changes in GNP. As a check for the robustness of the estimators, we also use a rolling window of the 12-month percentage change in the Business Cycle Index (BCI) current indicator as a measure of economic conditions. GNP and BCI data are obtained from the U.S. Department of Commerce, Bureau of the Census.

## V. Empirical Results

This Section reports the results of estimating the log-likelihood function in equation (9). The vector of explanatory variables $X_i(t_i)$ comprises the time-varying change in GNP, three 0-1 indicator variables to denote the rating (BB, B, and Not Rated), two 0-1 indicator variables to denote seniority (senior subordinated and senior), the coupon rate, a 0-1 indicator variable to identify bonds underwritten by Drexel, a 0-1 indicator variable that is equal to 1 for LBO-related bonds and zero otherwise, two 0-1 indicator variables to denote the years to maturity at issuance (between 10 and 15 years; greater than 15 years), and two 0-1 indicator variables to denote the issuance period (between 1980 and 1984; 1985 and subsequent). Issue size is included as a control variable.

The base group consists of high-yield bonds with the following properties: (1) subordinated, (2) rated CCC at issuance, (3) not underwritten by Drexel, (4) no coupon, (5) not LBO-related, (6) with initial maturity less than 10 years, and (7) issued prior to 1980. For both models, the base group assumes no change in GNP. Empirical results are presented in Table 3. Model I in Table 3 accounts for unobserved heterogeneity in the data, while model II does not. All estimates, along with the *t*-statistics, should be interpreted relative to the base group. The expected sign of the coefficients, based on the discussion in Section II, is denoted next to the variable name in the first column of Table 3.

***Aging Effect***

When applying a piece-wise approximation to an underlying continuous baseline hazard, one must determine the number and location of the discrete shifts. Working with monthly data with a maximum of 213 months, it clearly does not make sense to allow for a different $c$-parameter in every month. Instead, we chose to allow for a discrete shift every 24 months, with a last shift after 144 months.[13] Put differently, our model assumes constant monthly default rates between two shifts (e.g., $c_{25}=c_{26}=...=c_{48}$) and the $c$-coefficients indicate a proportional shift in the default rate relative to the first 24 months after issuance. Based on the $c$-estimates of model I, the baseline hazard is plotted in Figure 1.

There is strong evidence of an aging effect in the default behavior of the high-yield bonds. Every bi-annual jump, except for the last one, is positive and significant at the 1%

level, indicating that the monthly default rate increases significantly after the first two years. The monthly default rate for each of the first 24 months is estimated at 0.1% for the base group. The monthly default rate jumps to 0.46% [ = 0.001*exp(1.526)] after 24 months and 0.75% [ = 0.001*exp(2.015)] after 48 months. For the other bonds, the hazard is proportional to this graph, and is obtained by multiplying the baseline hazard with exp[$\beta X$], where $X$ is their vector of explanatory variables. For example, consider a bond with the following characteristics: rated CCC at issuance, junior, 13% coupon, $110 million issuance size, underwritten by Drexel, not LBO-related, initial maturity less than 10 years, issued after 1984. When the time-varying percentage change in GNP is zero, the monthly default rate in the first 24 months is 1.4% [ = 0.001*exp{11.492(0.13)-0.051(1.10)+0.289+0.880} = 0.001*exp{2.6069}] and 6.4% [ = exp{1.526}*0.014] in the second 24 months. After two years, the monthly default rate is 10.5% [ = exp{2.015}*0.014]. For the same type of bond, suppose now that the time-varying percentage change in GNP is 0.04%. In this case, the monthly default rates fall to 0.8% [ = 0.001*exp{2.6069-0.143(4)}], 3.7% and 6%, respectively.

The default rate continues to increase until the end of year 12, after which it seems to decline. However, a likelihood ratio test reveals that $c_{145...213}$ is not significantly different from $c_{121...144}$ ($X^2_{(1)}$=0.5227), thereby suggesting that the default rate levels off, rather than drops, after 12 years.[14]

Figure 1 also presents the baseline hazard of model II, which is the same as model I except that model II does not include the gamma mixing distribution; i.e., model II does not take unobserved heterogeneity into account. Figure 1 clearly illustrates the downward bias on the estimated aging effect when not accounting for unobserved heterogeneity: the baseline hazard for model II lies below the baseline hazard for model I. Put differently, if one does not explicitly correct for the fact that not all relevant factors can be included into the model, the aging effect will be underestimated. Note also that the effects of the covariates are larger and more significant in model I than in model II. Parameterizing unobserved heterogeneity therefore seems to have eliminated some of the attenuating effects of the omitted variables [see also Vanhuele et al. (1995) for a more detailed discussion].

***Economic Conditions***

It is important to note that we find a significant aging effect in both models even though we explicitly account for changing economic conditions. The coefficient estimate for the growth in GNP is negative and significant at the 1% level. This is consistent with significantly lower default rates when GNP increases. In other words, *the aging effect reported in prior studies is not solely a result of changing economic conditions*. Instead, our results suggest that both the number of months since issuance *and* the state of the economy have a significant impact on the default rate of high-yield bonds.

Model III in Table 3 provides estimates of the model when using the BCI as an alternative measure of the state of the economy. Again, the coefficient estimate is negative and significant. The similarity of the *c*-coefficients to those estimated when using GNP (model I) suggests that the aging effect is robust to different operationalizations of economic conditions.

***Issue-Specific Characteristics***

Model I in Table 3 also includes issue-specific variables. After controlling for the aging effect and changing economic conditions, BB rated bonds have a significantly lower default rate than CCC rated bonds. Specifically, the default rate for BB bonds is 75.8% [ =

100*(exp(-1.419)-1)] lower compared to CCC bonds. The default rates for B rated and non-rated bonds, however, are not significantly different than for CCC bonds.[15] Unlike prior studies [Altman (1989), Asquith et al. (1989) and Platt (1993)] we find that issues underwritten by Drexel do not have a significantly lower default rate.

Unlike Cheung et al. (1992), we find that the seniority of the debt does not have a significant impact on the default rate.[16] Similar to Rosengren (1993), issue size is not significant, while, as expected, the coupon rate has a positive and significant influence on the default probability.[17] Our results further indicate that LBO-related high-yield bonds are not significantly more likely to default than non-LBO related junk bonds.

We find that high-yield bonds issued in later periods (1980-1984 and subsequent to 1984) are more likely to default than those issued in the base period (1977-1979) at the 10% level. The relative magnitude of the corresponding parameter estimates (0.741 and 0.888) seems to provide some support for a further decline in bond issuance quality in the mid 1980's; however, the significance is not statistically different ($X^2_{(1)}$=0.1438).

Contrary to Cotter and Peck (1995), we find that bonds with the longest maturity do not have a significantly lower default rate than bonds with initial maturity less than ten years. This may be explained in part by the presence of some correlation between maturity and issuance year. As seen in Table 1, the maturity of debt at issuance declined over the sample period, which suggests a negative correlation between issuance period and maturity. The correlation between maturity and issuance year is -0.51 in our sample. While still below the conventional cut-off rule of 0.8 [Judge et al. (1988, p.868)], this may have affected to some extent the corresponding parameter estimates. We elaborate on this issue in Table 4, which provides estimates of models for subsets of the variables.

Model IV in Table 4 excludes the issuance period variables. For this model there is a negative and significant coefficient for the indicator variable for the longest maturity (greater than 15 years). Model V excludes the two maturity indicator variables. For model V, the coefficients of the two issuance period indicator variables are positive and significant, which is consistent with prior evidence of a decline in quality of bond issuances over the period. Note we still do not find a statistically significant difference between the coefficient for the 1980-1984 issuance period and the post-1984 issuance period ($X^2_{(1)}$=0.459).

The differing estimates of the maturity and issuance period coefficients between model I and models IV and V illustrate the difficulty associated with making statistical inferences in the presence of correlation between explanatory variables. One must question whether issuance period or maturity is the driving force behind prior research estimates. Indeed, a likelihood ratio test ($X^2_{(4)}$ = 6.567) indicates that we cannot reject the hypothesis that all four maturity and issuance period coefficients are equal to zero.

## VI. Conclusions

This empirical study investigates the individual-level default rate of high-yield corporate bonds. Previous studies of high-yield bond defaults provided evidence of an aging effect: the longer bonds have been outstanding, the greater the default probability. However Blume, Keim and Patel (1991) asserted that this aging relationship reflects changing economic conditions. This study of bond defaults discriminates between the aging effect and the impact of macroeconomic changes.

We adopt a hazard model where the impact of aging is modeled semi-parametrically, which is flexible enough to capture monotonic or non-monotonic aging effects. Our specification includes both fixed and time-varying variables, which allows us to estimate the impact of aging, changing economic conditions and issue-specific characteristics on the default rate simultaneously. The specification also takes into account the presence of unobserved heterogeneity.

The model is applied to a sample of 579 high-yield non-convertible bonds issued between 1977 and 1989 and followed through December 1994. We find a significant aging effect even though we explicitly account for changing economic conditions. Our findings suggest that default rates exhibit a positive aging effect. In addition, we find that the rating at issuance, the size of the coupon and the issuance period are significantly related to the default rate. However, the issue's seniority, size, maturity, whether the issue was underwritten by Drexel and whether the issue was LBO-related do not appear to have a significant influence on the likelihood of default.

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TABLE 1 - EXPLANATORY VARIABLES BY YEAR OF ISSUANCE

| Issue Year | Issues | Average Size ($Mil.) | Mean Coupon | Mean Time to Maturity (in years) | Rated BB | Rated B | Rated CCC | Not Rated (NR) | Defaults |
|---|---|---|---|---|---|---|---|---|---|
| 77 | 19 | 30.158 | 10.47 | 17.21 | 10 | 8 | 1 | 0 | 3 |
| | | | | | 52.6% | 42.1% | 5.3% | 0.0% | 15.8% |
| 78 | 33 | 27.197 | 11.54 | 18.64 | 5 | 27 | 1 | 0 | 11 |
| | | | | | 15.2% | 81.8% | 3.0% | 0.0% | 33.3% |
| 79 | 27 | 26.611 | 12.34 | 18.07 | 3 | 21 | 1 | 2 | 10 |
| | | | | | 11.1% | 77.8% | 3.7% | 7.4% | 37.0% |
| 80 | 28 | 32.475 | 13.74 | 18.29 | 5 | 20 | 1 | 2 | 7 |
| | | | | | 17.9% | 71.4% | 3.6% | 7.1% | 25.0% |
| 81 | 13 | 65.769 | 13.67 | 17.00 | 3 | 9 | 0 | 1 | 7 |
| | | | | | 23.1% | 69.2% | 0.0% | 7.7% | 53.8% |
| 82 | 26 | 61.538 | 13.39 | 14.31 | 6 | 18 | 0 | 2 | 9 |
| | | | | | 23.1% | 69.2% | 0.0% | 7.7% | 34.6% |
| 83 | 53 | 95.292 | 11.98 | 15.68 | 16 | 25 | 4 | 8 | 23 |
| | | | | | 30.2% | 47.2% | 7.5% | 15.1% | 43.4% |
| 84 | 52 | 68.644 | 13.61 | 12.56 | 6 | 27 | 4 | 15 | 13 |
| | | | | | 11.6% | 51.9% | 7.7% | 28.8% | 25.0% |
| 85 | 72 | 90.972 | 12.98 | 11.54 | 10 | 42 | 4 | 16 | 15 |
| | | | | | 13.9% | 58.3% | 5.6% | 22.2% | 20.8% |
| 86 | 90 | 132.644 | 11.69 | 11.84 | 12 | 57 | 11 | 10 | 22 |
| | | | | | 13.3% | 63.3% | 12.2% | 11.1% | 24.4% |
| 87 | 76 | 182.086 | 12.75 | 11.13 | 14 | 47 | 14 | 1 | 22 |
| | | | | | 18.4% | 61.8% | 18.3% | 1.3% | 28.9% |
| 88 | 49 | 202.806 | 13.27 | 10.63 | 5 | 31 | 11 | 2 | 10 |
| | | | | | 10.2% | 63.3% | 22.4% | 4.1% | 20.4% |
| 89 | 41 | 202.049 | 13.29 | 11.63 | 5 | 30 | 3 | 3 | 8 |
| | | | | | 12.2% | 73.2% | 7.3% | 7.3% | 19.5% |
| All | 579 | 111.781 | 12.64 | 13.40 | 100 | 362 | 55 | 62 | 160 |
| | | | | | 17.3% | 62.5% | 9.5% | 10.7% | 27.6% |

Notes: The sample consists of 579 high-yield securities issued between 1977 and 1989, for which data is available from *Compustat* and the Standard and Poor's *Monthly Bond Guides.* The amount issued, number of issues, mean issue amount, mean coupon, mean number of years from issuance to maturity, are from Securities Data Corporation. Bond status, as of December 31, 1994, is determined using the Standard and Poor's *Monthly Bond Guides*, Moody's *Bond Record*, the *Wall Street Journal*, and *Lexis/Nexis* wire service reports.

The percentages of the bonds rated BB, B, CCC and Not Rated in a given year sum to 100%.

TABLE 2 - ORIGINAL RATING, SENIORITY, UNDERWRITER (DREXEL) AND PURPOSE (LBO-RELATED)

| | Issues | Defaults | Mean Time to Default (in years) | Mean Time to Censored (in years) |
|---|---|---|---|---|
| Original Rating | | | | |
| BB | 100 | 20<br>20.0% | 5.97 | 6.99 |
| B | 362 | 105<br>29.0% | 4.54 | 6.80 |
| CCC | 55 | 16<br>29.1% | 3.29 | 6.16 |
| Not rated | 62 | 19<br>30.6% | 4.10 | 7.51 |
| Seniority | | | | |
| Mortgage | 18 | 5<br>27.8% | 5.22 | 5.48 |
| Senior | 103 | 27<br>26.2% | 4.51 | 6.23 |
| Senior Subordinated | 227 | 60<br>26.4% | 4.00 | 6.35 |
| Subordinated | 231 | 68<br>29.4% | 4.99 | 7.77 |
| Underwritten by Drexel | 230 | 70<br>30.4% | 4.54 | 6.62 |
| Leveraged Buyout | 60 | 14<br>23.3% | 3.02 | 5.76 |
| All | 579 | 160<br>27.6% | 4.54 | 6.85 |

Notes: Sources of data are described in the notes for Table 1. The bond seniority and underwriter are from the Standard and Poor's *Monthly Bond Guides*. Bonds associated with LBOs were identified by Securities Data Corporation.

TABLE 3 - PARAMETER ESTIMATES OF THE HAZARD MODEL

| Variable | | I With gamma mixing distribution | II Without gamma mixing distribution | III With BCI with gamma mixing distribution |
|---|---|---|---|---|
| Baseline Hazard | | | | |
| r/a (I,III) or $h_0$ (II) (1-24 months) | | 0.001 | 0.001 | 0.001 |
| $c_{25..48}$ | | 1.526[a] | 1.145[a] | 1.514[a] |
| $c_{49..72}$ | | 2.015[a] | 1.121[a] | 1.983[a] |
| $c_{73..96}$ | | 2.883[a] | 1.354[a] | 2.826[a] |
| $c_{97..120}$ | | 3.192[a] | 1.178[a] | 3.122[a] |
| $c_{121..144}$ | | 3.341[a] | 1.041[c] | 3.253[a] |
| $c_{145..213}$ | | 2.517[c] | 0.071 | 2.426[c] |
| Economic Conditions | | | | |
| Change in GNP (I,II) | (-) | -0.143[a] | -0.091[b] | |
| Change in BCI (III) | (-) | | | -0.126[a] |
| Issue-specific Variables | | | | |
| BB rating | (-) | -1.419[b] | -0.515[c] | -1.390[b] |
| B rating | (-) | -0.398 | -0.054 | -0.382 |
| Not rated | (+) | -0.304 | -0.185 | -0.297 |
| Senior debt | (-) | 0.209 | 0.180 | 0.202 |
| Senior subord. debt | (-) | -0.028 | -0.029 | -0.034 |
| Coupon rate | (+) | 11.492[b] | 6.124[c] | 11.458[b] |
| Issue size | | -0.051 | -0.039 | -0.052 |
| Drexel | (-) | 0.289 | 0.199 | 0.293[c] |
| LBO | (+) | 0.118 | -0.056 | 0.114 |
| Maturity: | | | | |
| 10 to 15 years | (-) | -0.280 | -0.214 | -0.279 |
| Greater than 15 years | (-) | -0.516 | -0.359 | -0.513 |
| Issuance Period: | | | | |
| 1980 - 1984 | (+) | 0.741[b] | 0.295 | 0.701[b] |
| After 1984 | (+) | 0.880[b] | 0.308 | 0.821[b] |
| Log-likelihood | | -1024.824 | -1027.509 | -1025.201 |

Notes: [a(b,c)] indicates significant at the 1% (5%,10%) level for a two-tailed test (the baseline hazard) or a one-tailed test (the other explanatory variables), based on asymptotic *t*-statistics.

Expected coefficient signs appear between parentheses next to the variable names.

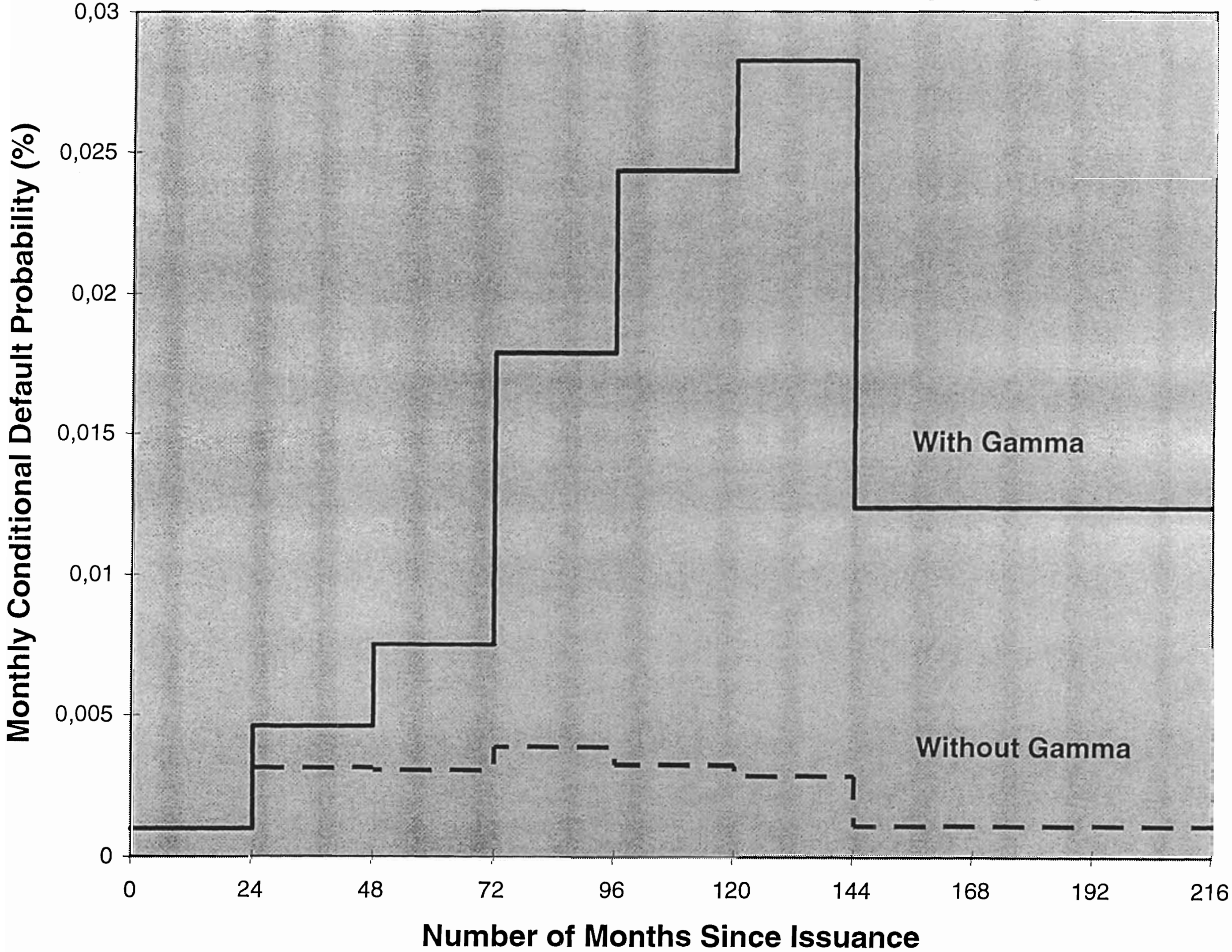
TABLE 4 - PARAMETER ESTIMATES OF THE HAZARD MODEL WITH THE GAMMA MIXING DISTRIBUTION EXCLUDING ISSUANCE PERIOD AND MATURITY VARIABLES

| | | I | IV | V |
|---|---|---|---|---|
| | | All explanatory variables | Excluding issuance period variables | Excluding maturity variables |
| Baseline Hazard | | | | |
| r/a (1-24 months) | | 0.001 | 0.001 | 0.000 |
| $c_{25...48}$ | | 1.526[a] | 1.476[a] | 1.530[a] |
| $c_{49...72}$ | | 2.015[a] | 1.905[a] | 2.024[a] |
| $c_{73...96}$ | | 2.883[a] | 2.669[a] | 2.896[a] |
| $c_{97...120}$ | | 3.192[a] | 2.887[a] | 3.210[a] |
| $c_{121...144}$ | | 3.341[a] | 2.978[a] | 3.363[a] |
| $c_{145...213}$ | | 2.517[c] | 2.126 | 2.543[c] |
| Economic Conditions | | | | |
| Change in GNP | (-) | -0.143[a] | -0.128[a] | -0.143[a] |
| Issue-specific Variables | | | | |
| BB rating | (-) | -1.419[b] | -1.345[b] | -1.436[a] |
| B rating | (-) | -0.398 | -0.384 | -0.393 |
| Not rated | (+) | -0.304 | -0.269 | -0.240 |
| Senior debt | (-) | 0.209 | 0.024 | 0.264 |
| Senior subord. debt | (-) | 0.028 | 0.014 | -0.006 |
| Coupon rate | (+) | 11.492[b] | 13.118[a] | 9.746[b] |
| Issue size | | -0.051 | -0.002 | -0.060 |
| Drexel | (-) | 0.289 | 0.310 | 0.331[c] |
| LBO | (+) | 0.118 | 0.150 | 0.131 |
| Maturity: | | | | |
| 10 to 15 years | (-) | -0.280 | -0.344 | |
| Greater than 15 years | (-) | -0.516 | -0.772[b] | |
| Issuance Period | | | | |
| 1980 - 1984 | (+) | 0.741[b] | | 0.888[a] |
| After 1984 | (+) | 0.880[b] | | 1.122[a] |
| Log-likelihood | | -1024.824 | -1026.312 | -1025.362 |

Notes: [a(b,c)] indicates significant at the 1% (5%,10%) level for a two-tailed test (the baseline hazard) or a one-tailed test (the other explanatory variables), based on asymptotic *t*-statistics., based on asymptotic *t*-statistics.

Expected coefficient signs appear between parentheses next to the variable names.

**Figure 1**
**Time Dependence of the Conditional Default Probability of High-Yield Bonds**



ENDNOTES

1. Studies which provide evidence of an increase in default rates during the first few years since issuance include Altman (1989), Altman and Kishore (1995), Asquith, Mullins and Wolff (1989), Cheung, Bencivenga and Fabozzi (1992), and Platt (1993).

2. One exception is Rosengren (1993), who used a logit model. Section II compares logit models with our method of analysis.

3. In the actuarial approach, default rates are measured as the amount of defaults relative to the population of bonds in the same cohort that could still default (i.e., that were still at risk).

4. Blume and Keim (1991) and Cheung, Bencivenga and Fabozzi (1992) also documented the high number of defaults in some years relative to others. Jónsson and Fridson (1995) provided evidence that aggregate default rates are related to changes in economic conditions, as measured by the percentage change in Gross National Product.

5. Ignoring unobserved heterogeneity is conceptually similar to the omitted variable problem, and may lead to biased parameter estimates.

6. Examples of applications of this semi-parametric approach include Vanhuele, Dekimpe, Sharma and Morrison (1995) and Han and Hausman (1990).

7. To avoid identification problems when simultaneously estimating $c_1$ and $h_0$, no separate indicator variable is added for the first period.

8. Note that we assume the censoring occurs at the beginning of the grouping interval. Clearly, some such assumption is needed given the discrete nature of our data.

9. It should be noted that the contribution of a bond that is still outstanding, has been called, or was involved in a non-distressed exchange is given by $S(t_i\text{-}1)$, while the contribution of a bond that has matured is $S(t_i)$. This difference in contributions, due to the grouping intervals, does not necessitate a special term in the likelihood function if the input matrix is altered by adding 1 to the life span of every bond that has matured.

10. See Vanhuele et al. (1995) for a derivation.

11. We validated the default and call dates for corresponding bonds in the Blume and Keim (1991) Appendix III for those bonds which defaulted or were called during their analysis period.

12. The aggregate ratings include bonds rated with a + or -; e.g., the rating category B includes bonds rated B+ and B- at issuance.

13. No other jumps are allowed after 144 months so as to ensure that the *c*-estimates are based on a large enough group of bonds.

14. The aging effect is not likely to be affected by crisis at maturity: only one out of 579 bonds in our sample defaulted at maturity.

15. Since the ratings categories are broad rather than detailed categories, adjusting the ratings for seniority as in Fridson and Garman (1995) does not change the results.

16. Because of the small number of secured issues, mortgage issues are included in the senior issues classification.

17. All bonds in our sample have fixed coupons. Estimating the model with the coupon rate relative to the average coupon rate in the issuance year did not materially change the results.